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# A New Look at Family Migration and Women's Employment Status

*Family migration has a negative impact on women's employment status. Using longitudinal data from the British Household Panel Survey (3,617 women; 22,354 women/wave observations) we consider two neglected issues. First, instead of relying on the distance moved to distinguish employment-related migrations, we use information on the reason for moving, allowing us to separate employment-related moves, stimulated by the man or the woman, from other moves. Second, we consider selection effects and the role of state dependence in relation to women's employment status prior to moving. Moving for the sake of the man's job has a significant negative effect on subsequent employment status for previously employed women. Women who were not employed previously benefited only slightly from family migration.*

The influence of ''family migration,'' or the long-distance moves of partnered individuals, on employment status has attracted a considerable literature over the past 30 or so years. The general consensus is that families are more likely to move in support of the man's career and that the woman's employment status is likely to suffer as a result of such moves. According to human capital theory, families weigh the benefits of overall gains associated with a move on behalf of one person's career against the negative effects of disrupting the partner's employment (Becker, 1974). Although a genderless theory, moves tend to be made to support the man's career more often than the women's and, as a result, women are more likely to be ''trailing spouses'' or ''tied migrants.''

More recently, there have been some moves away from relatively crude human capital interpretations and many argue for a more nuanced approach that recognizes the influence of gendered family resources (e.g., Bielby & Bielby, 1992; Halfacree, 1995; Juerges, 1998, 2006; Shauman & Noonan, 2007; Shihadeh, 1991). Thus, Boyle, Cooke, Halfacree, and Smith (1999a) showed that women are more likely to be unemployed or economically inactive following long-distance family migration even in those cases when the women had a higher status occupation than their partner. This does not seem to accord with a gender-neutral human capital model. Also, we need to recognize that women may suffer from family moves even when the underlying reason was not employment focused. Even short-distance changes in residence can influence psychological well-being and depression, particularly among women who are

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often expected to cope with the practicalities of moving house (Brett, 1980; Makowsky, Cook, Berger, & Powell, 1988; Weissman & Paykel, 1972). These include arranging for the movement of household possessions, acquiring new household items, and, for those with children, organizing child care and other child-centered activities (Magdol, 2002). It may be that these types of effects will have an influence on women's desire or ability to work in the labor market, even following shorter-distance moves.

Numerous early studies confirmed that women's labor market status suffers as a result of family migration (e.g., Lichter, 1980, 1982; Long, 1974; Mincer, 1978; Morrison & Lichter, 1988; Sjaastad, 1962; Spitze, 1984) and, more recently, Boyle, Halfacree, and Smith (1999) used comparative, cross-national data for the United Kingdon and the United States at the beginning of the 1990s and showed that women continued to be more likely to be out of work following family migration (Boyle, Cooke, Halfacree, & Smith, 1999b, 2001). This result was remarkably consistent in the United Kingdom and the United States (Boyle, Cooke, Halfacree, & Smith, 2002), even controlling for motherhood status (Boyle, Cooke, Halfacree, & Smith, 2003) and the relative occupational status of the partners (Boyle et al., 1999a). Other recent studies confirm these broad conclusions (Bailey & Cooke, 1998; Cooke, 2001; Jacobsen & Levin, 1997, 2000; Shihadeh, 1991; Smits, 1999; Taylor, 2007), although some cast doubt over the importance of these findings. For example, Clark and Withers (2002) argue that although women's labor force participation is disrupted by family migration, these effects tend to be short-lived in the United States. Even so, Clark also found that the disruptive effects were longer term in Britain (Clark & Huang, 2006). Also, we should recognize that spells of unemployment have been found to have long-term ''scarring'' effects in a number of studies (Arulampalam, 2001; Arulampalam, Gregg, & Gregory, 2001; Eliason & Storrie, 2006; Gregg & Tominey, 2004) and a recent U.K./U.S. analysis shows that although the effects of family migration on women's employment status are not as great as the effect of having a child, they are considerable (Cooke, Boyle, Couch, & Feijten, in press).

This paper extends this body of research in two distinct ways. First, studies in the past have tended to use long-distance moves, or moves that cross an essentially arbitrary administrative boundary, as surrogate measures for employment-related family migration. Shorter distance moves are usually assumed to be more likely to be housing related. This approach is often forced upon researchers by the available information—census data do not include the reason for migration—but this inevitably misclassifies some moves, as not all long-distance moves are employment related, and some shorter distance moves are for work-related reasons. The extent of the bias that this introduces is not clear, and it is certainly the case that the results in previous family migration studies cannot be explained solely by moves stimulated by the man's career. We use longitudinal data drawn from the British Household Panel Survey (BHPS), which includes information about the reason for the move, and we compare this with the more common distance-based approach. In particular, we are able to separate employment-related moves, stimulated by the man, the woman, or both partners, from other types of moves. We are then able to assess whether *moving for the man's job* has a significant negative effect on women's employment status compared to being immobile or moving for other reasons. Only Taylor (2007) appears to have used the reason for moving in a recent study of family migration (although see also Clark & Withers, 2007).

Second, we consider the role of ''state dependence,'' which has been ignored in family migration studies to date. This refers to the likelihood that a person's characteristics are similar through time; in our example, we might assume that women who were out of employment at $t$-1 are also more likely to be out of work at time $t$ and this persistence needs to be controlled for. As Heckman (2001, p. 706) notes in his Nobel speech: ''A frequently noted empirical regularity in the analysis of unemployment data is that those who were unemployed in the past or have worked in the past are more likely to be unemployed (or working) in the future.''

This regularity has been established in a wide literature (e.g., Heckman, 1980, 1991; Hyslop, 1999; Muhleisen & Zimmermann, 1994; van den Berg & van Ours, 1996), including in a number of British studies (e.g., Arulampalam, Booth, & Taylor, 2000; Davies, Elias, & Penn, 1992; Gregg, 2001, Narendranathan & Elias, 1993). From a human capital perspective, state dependence in unemployment occurs because of the deterioration of existing human capital during an unemployment spell as well as from the nonaccumulation of new human capital during this period. Potential employers may also use previous labor market history as a signal of productivity (Eliason & Storrie, 2006), and there is evidence that unemployed people are more likely to become reemployed in low-wage, poor quality jobs that raise the risk of subsequent unemployment (Stewart, 2005). Ignoring this issue may have important implications in studies of family migration, as women who were previously out of work may be more willing to participate in family migration than those who were employed. Rather than migration reducing the likelihood that a women works, women who were previously out of work may be more likely to engage in family migration. This potentially important selection effect has been ignored in previous family migration studies and could have resulted in biased models and underestimated standard errors. Thus, we fit appropriate dynamic panel models that adjust for unobserved heterogeneity and initial conditions to account for state dependence.

In combination, this study provides a unique insight into the influence of family migration on women's employment status, using up-to-date longitudinal data that allow us to explore transitions in and out of different states. Using the reason for the move allows a direct test of the family migration hypothesis, which has rarely been achieved in the past, and the control introduced for state dependence is an important analytical improvement, given the strength of state dependence known to exist in (un)employment. In the remainder of the paper we describe the data, the regression methods that we use, and the results before discussing the implications of our findings.

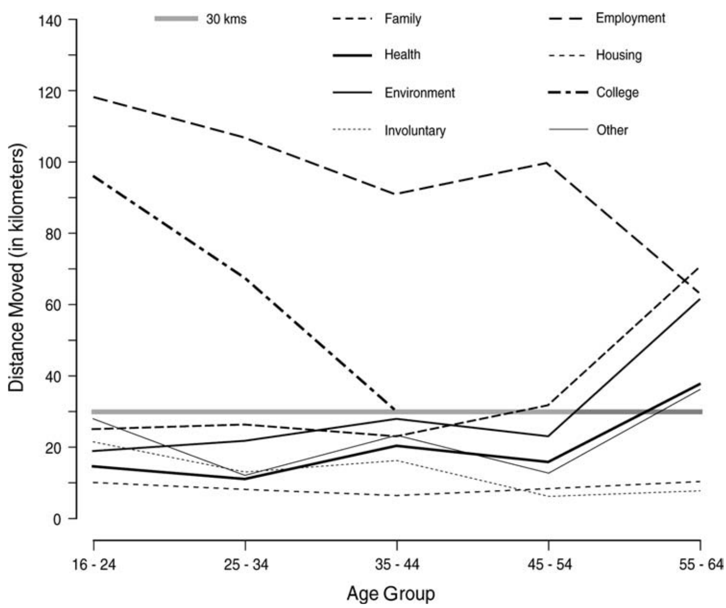
## METHOD

### *Data*

We utilized the British Household Panel Study, which has been collecting data on a nationally representative sample of households since 1991. Interviews are carried out on household members aged 16 and over every year, and respondents provide information in regard to current situations and changes since the previous year. We extracted information from Waves 2 – 12 (B – N) on a restricted sample of 4,491 women aged 16 – 64 in married or cohabiting partnerships, which gives us 29,349 woman/wave observations. Women may have joined the panel since 1992 or could have left the panel before 2002. Those whose relationship ended were dropped from the sample, but they returned if they began a new partnership. In total, 1,043 women were captured in all 11 waves and answered all the relevant questions accounting for 23.2% of all women in this sample. Nearly half (45%) of the women appeared in two or more waves continuously, 13% appeared only once, and 12% were missing for one or more waves but returned to the panel at a later stage. Of the 4,491 women only 3,617 answered all the relevant questions, and hence these were the group included in the modeling. Data were extracted for a range of individual- and area-level variables.

Our outcome variable compares women who were unemployed or economically inactive (about one third of the observations) with those who were in employment. We control for a range of explanatory variables, but focus particularly on migration status, for which two variables were used. The first compares those who moved 30 km or more between $t - 1$ and $t$ to those who moved only short distances or who did not move at all. Thirty kilometers is the average distance moved by couples in the BHPS (Böheim & Taylor, 2002). The use of a distance variable is a common strategy used in the absence of further information on the reason for the move to distinguish probable employment-related moves from moves stimulated by housing or other reasons. Fortunately, though, the BHPS includes data on the reason for moving (Taylor, 2007), and we therefore constructed a second variable that compares nonmovers with those who moved for various reasons. Among all long distance moves, 46.2% were for employment reasons. Of those moves for employment reasons, 82.6% were over a long distance. Figure 1 plots the average distance moved by age and the reason for the move from the BHPS data. It is clear that, on average, employment-related moves tend to be longer than 30 km (the distance cutoff we have chosen). Moves into college for younger adults and moves for family and environmental reasons for older adults also tend to be longer than 30 km, however, demonstrating the inadequacy of assuming that an arbitrary distance cutoff distinguishes between employment-related and other types of moves. Consequently, we also used the reason for the move to divide migrants into those who moved for the man's job, those who moved for either the woman's or both jobs, and those who moved for other reasons. Around 15% of all moves were for at least one partner's job. In line with the family migration literature, moving for the man's job is hypothesized to have

FIGURE 1. AVERAGE MIGRATION DISTANCE FOR WOMEN BY AGE AND REASON FOR THE MOVE, ACROSS ALL WAVES.




*Note:* Bold horizontal line represents 30 km. This figure includes multiple moves made by the same individuals in different waves.

a negative impact on women's employment status, whereas we would expect that women who moved for either their job or for both jobs should be more likely to be employed following the move.

It is also important to consider possible state dependence in (un)employment characteristics—an issue that seems to have been ignored to date in the family migration literature. The assumption is that women who were out of employment at $t - 1$ are also more likely to be out of work at time $t$ and this persistence needs to be controlled for. Women who are out of work may be more willing to engage in family migration than women who are working, and it is therefore important to test whether women who were employed at $t - 1$ were affected by family migration. Thus, we included a 1-year lag of employment status in our models but are mindful that the inclusion of such a variable is likely to lead to bias in conventional random effects panel models. Hence we adopted a dynamic modeling approach that addresses this problem (see below).

In addition, we included a range of individual, household, and contextual time-varying variables expected to influence employment status (Table 1). Individual-level variables included basic demographic controls of age, qualifications, and marital status. Household variables included car ownership (women in households with one or two or more cars were expected to be less likely to be out of work following family migration, because job access may not have been such a problem), the life course stage of the family measured by the age of youngest child (women with young children are expected to be less likely to be in the labor market), household size (women in larger households were also expected to be less likely to be in the labor market), and the comparative occupational status of the woman compared to her partner. This was measured using the Cambridge score, which is a gender-specific occupation-based measure of social stratification (Prandy, 2000), constructed by comparing the occupations of married partners and friends. We might expect women to suffer more from family migration when the occupational gap between her and her partner's occupation was large. We also considered the geographical context in which family migration is played out. We might imagine that certain

Table 1. *Descriptive Statistics for Variables Across All Waves (N = 29,349)*

| Variable | Mean |
|---|---|
| Employment status | |
| Employed | 0.676 |
| Not employed | 0.324 |
| Migration | |
| Nonmover/short distance mover | 0.986 |
| Long distance mover (>30 km) | 0.014 |
| Age group | |
| 16 – 24 | 0.073 |
| 25 – 34 | 0.267 |
| 35 – 44 | 0.268 |
| 45 – 54 | 0.240 |
| 55 – 64 | 0.153 |
| Qualifications | |
| No qualification | 0.292 |
| Intermediate qualification | 0.598 |
| Degree and higher | 0.110 |
| Marital status | |
| Married | 0.819 |
| Cohabiting | 0.181 |
| Car ownership | |
| No car | 0.097 |
| 1 car | 0.449 |
| 2+ cars | 0.454 |
| Youngest child | |
| No child | 0.447 |
| 0 – 4 years | 0.199 |
| 5 – 10 years | 0.128 |
| 11 – 15 years | 0.088 |
| Nondependent children | 0.138 |
| Household size | |
| 2 and less people | 0.373 |
| 3 people | 0.499 |
| 4 or more people | 0.129 |
| Region | |
| London | 0.084 |
| Other metropolitan | 0.173 |
| Rest of South East | 0.422 |
| Rest of South and North | 0.321 |
| Employment rate | |
| Women's rate < 75% percentile men's | 0.758 |
| Women's rate ≥ 75% percentile men's | 0.242 |
| Cambridge score | |
| Women's Cambridge score ≤ men's | 0.573 |
| Women's Cambridge score > men's | 0.427 |

*Note:* Qualifications $N = 28{,}546$, Cambridge score $N = 22{,}864$.

locations provide better opportunities for women to find employment following a move than others, and we contrasted London with the other metropolitan areas, the South East remainder, and the rest of Southern and Northern Britain with the expectation that women are more likely to fi labor market employment in the large urban centers. We consider this issue more directly by including a measure of women's employment rates in the Local Authority District in which they lived to test whether women are more likely to be in work in areas where women's employment rates are high compared to the national average.

### *Procedure*

We have panel data with repeated observations for the same partnered woman, and our outcome measure is a binary variable distinguishing those who are, or are not, in employment at time *t*. We therefore used random effects probit models, appropriate for modeling such panel data. We chose a random effects *dynamic* model because no equivalent fixed effects model is available. We were particularly interested in whether employment status in the previous wave ($t-1$) had an influence on the outcome at time *t* (state dependence), but the inclusion of a lagged *y* variable is known to introduce bias. A naïve model that simply includes a lagged *y* variable could overestimate the impact of state dependence. We therefore fitted Heckman-type (1981) random-effects dynamic probit models that control for initial conditions using an approach developed by Stewart (2006). The model can be written as

$$y^*_{it} = \gamma y_{it-1} + x'_{it}\beta + \alpha_i + u_{it} \quad (i = 1, \ldots, N;\ t = 2, \ldots, T); \qquad (1)$$

where $y^*_{it}$ is the unobservable propensity of the wife to be out of employment at time $t$, $x_{it}$ is a vector of explanatory variables, $\alpha_i$ are (unobserved) individual-specific random effects, and $u_{it}$ are assumed to be distributed $N(0, \sigma^2_u)$. The estimation of this model requires an assumption about the initial conditions, $y_{i1}$ and their relationship with $\alpha_i$. The initial conditions can be assumed to be exogenous if the start of the process coincided with the beginning of the observation period for each person, but this is not the case in these panel data. If the initial conditions are correlated with $\alpha_i$, we would expect $\gamma$ to be overestimated, leading to an overstated impact of state dependence (Chay & Hyslop, 2000).

We therefore adopted an approach to the initial conditions problem that involves a linearized reduced form equation for the initial period:

$$y_{i1} = z'_{i1}\pi + \eta_i \quad (i = 1, \ldots, N), \qquad (2)$$

where $z_{i1}$ is a vector of exogenous instruments that includes $x_{i1}$, and $\eta_i$ is correlated with $\alpha_i$, but uncorrelated with $u_{it}$ for $t \geq 2$. In our models, our instrument variables, which were significant in a simple probit model fitted for $t = 1$, but insignificant in a probit model for $t \geq 2$, were household type (couple with no children, dependent children, or nondependent children), occupational diversity (a measure of the variety of occupations in which women were employed in the Local Authority District of residence, calculated using Shannon's entropy, which is a measure of diversity; Shannon, 1948) and occupational penetration (the success of women in men's occupations in the Local Authority District, calculated using a dissimilarity index). Below, we compare the results from the standard and the dynamic random effects models. We also calculate the probabilities of employment/economic inactivity for different migrant subgroups and, to make these comparable to those that would be derived from standard probit estimates, we multiplied by an estimate of $\sqrt{1-\lambda}$ where $\lambda$ is the ratio of the heterogeneity variance to the total variance.

Because it is possible that attrition bias (Goodman & Blum, 1996) may have influenced our results, we also conducted a sensitivity analysis that involved fitting the same set of models for the 1,043 women in our sample who appeared in all waves and answered all the relevant questions. If the results from this balanced sample were broadly similar, it suggests that attrition bias does not influence the broad conclusions of the analysis.

## RESULTS

Table 2 demonstrates that there is considerable state dependence in unemployment/economic inactivity in the raw data. Focusing on the group of nonmigrant/short-distance migrant women, 91% of those who were employed at $t - 1$ were also employed at $t$, whereas 82% of women who were not employed at $t - 1$ were also not employed at $t$. We also see, however, that long-distance migration reduces women's employment rates. Although only 9% of previously employed women who did not move over 30 km were unemployed or economically inactive at time $t$, 27% of previously employed women who moved a long distance were out of work at time $t$. On the other hand, the comparable figures for women who were not employed at $t - 1$ were 82% and 73%. In the raw data, at least, moving appears to reduce the likelihood of employment for women who were previously employed, but it increases the likelihood that a woman is employed if she was out of work prior to the move. This demonstrates the necessity for controlling for state dependence in our modeling analysis. Family migration may not have negative labor market effects for all women—it may even have positive effects for some.

Table 3 compares the distance moved by the reason for the move. This demonstrates that

Table 3. *Distance Moved by the Reason for the Move, All Movers in All Waves (row %)*

| Migration Status | Distance Moved: <30 km | Distance Moved: ≥30 km | Total Observations |
|---|---|---|---|
| Moved woman's or both jobs | 60 | 40 | 208 |
| Moved man's job | 50 | 50 | 212 |
| Moved other reasons | 91 | 9 | 2,464 |

Table 2. *Employment Status at Time* t *by Migration Status/Employment Status at* t *− 1 (row %)*

| Migration Status | Employment Status ($t$ - 1) | Employment Status ($t$): Employed | Employment Status ($t$): Not Employed | Total Observations |
|---|---|---|---|---|
| Nonmigrant/short distance migrant | Employed | 91 | 9 | 16,777 |
| | Not employed | 18 | 82 | 7,711 |
| Long distance migrant (30 km+) | Employed | 73 | 27 | 227 |
| | Not employed | 27 | 73 | 143 |

a considerable proportion of moves that occur for employment reasons occurred over less than 30 km. This is particularly the case for moves for the woman's or both jobs, as 60% of these were less than 30 km, whereas 9% of moves for other reasons were 30 km or more. The use of a simple distance cutoff, as in most previous research, clearly misallocates quite a high proportion of moves—most of the employment-related moves are actually over short distances, even though our choice of cutoff is relatively short at 30 km.

Table 4 presents the results from three models, two of which are standard random-effects probit models and the third is a random-effects dynamic probit model. All three models use the distance-based definition of migration. Figure 2a – c provides the calculated probabilities for different population subgroups for Models 1 – 3.

Model 1 (Table 4, Figure 2a) shows that women who moved long distances were significantly more likely to be unemployed or economically inactive than nonmigrants or those who moved short distances. This confirms the findings from previous studies. Model 2 (Table 4) includes the interaction with the lagged employment variable. The probabilities for women by migrant status and lagged employment status are displayed in Figure 2b. This is a standard random-effects probit, however, and, as described above, it is likely to be biased because of the incorporation of the lagged *y* variable. Model 3 (Table 4, Figure 2c) includes the same set of variables as Model 2, but it is a random-effects *dynamic* probit. The results show that being out of work at $t - 1$ was strongly and positively related to being out of work at time $t$. The migration variable remained positive and significant, but has increased in size compared to Model 1, and the interaction between migration and the lagged employment variable is negative and significant. Women who were employed at $t - 1$ and move a long distance have over twice the probability of being out of work than similar women who did not move a long distance; this difference was considerably greater than in Figure 2a, when we did not distinguish prior employment status. The highest probability of unemployment/economic inactivity, however, was for women who were out of work at $t - 1$ and who did not move a long distance. In line with the results for the raw data (Table 3), migrant women who were not employed at $t - 1$ were less likely to be out of work than unemployed/economically inactive women at $t - 1$ who did not migrate a long distance (although the difference between these two groups was not statistically significant).

Comparing Models 2 and 3 (Table 4, Figures 2b, 2c), and as expected, we find that the state dependence effect was overstated in Model 2. The parameter for the lagged *y* variable reduces slightly in Model 3, as does the parameter for the interaction between this variable and migration. In both cases the standard errors increased, although both variables remained highly signifi nt. Overall, though, the broad conclusions are similar for both models: Long distance migration has a signifi t negative effect on those women who were in employment at $t - 1$.

Briefl the results for the other explanatory variables are consistent with theoretical expectations. Those likely to be out of work included the oldest group (ages 55 – 64) and those with young children. Women who were more likely to be working included those with higher qualifi tions, who owned one or more cars, with nondependent children, living in areas where women's employment rates were high, and with a higher male Cambridge Scale score than their partner.

Table 5 provides similar comparative results, except that we use the reason for the move, rather than the distance moved. Model 4 ignores state dependence and shows that although moving for the woman's or both jobs increased the chance of working, moving for either the man's job or for other reasons decreased the chance of working (Figure 3a). Model 5 (Table 5, Figure 3b) includes the lagged employment variable and its interaction with the reason for moving. Model 6 (Table 5, Figure 3c) is the dynamic random-effects probit model. This shows that being out of work at $t - 1$ increased the probability of being out of work at time $t$, although this effect was attenuated if the family moved. For those who were employed at $t - 1$, moving increased the probability of the woman being out of work at time $t$, whatever the reason for the move. Even moving for the woman's or both jobs increased the probability of unemployment/economic inactivity, although not significantly compared to nonmigrants. If the move was stimulated by the man's job, the probability of being out of work is raised considerably to over twice that for nonmigrant women. Moving for other reasons also increased the probability of being out of work significantly.

Table 4. *Model Parameters for Woman's Employment Status (y ¼ 1, Not Employed) Using Distance Moved*

| Categories | Model 1, Random-Effects Probit | | Model 2, Random-Effects Probit | | Model 3, Random-Effects Dynamic Probit[a] | |
|---|---|---|---|---|---|---|
| | Parameters | *SE* | Parameters | *SE* | Parameters | *SE* |
| Employment status ($t$-1) | | | | | | |
| Employed (base) | | | | | | |
| Not employed | | | 1.626* | 0.037 | 1.297* | 0.043 |
| Migration (30 km1) | | | | | | |
| Nonmigrant/short distance migrant (base) | | | | | | |
| Long distance mover | 0.548* | 0.102 | 0.879* | 0.107 | 0.968* | 0.119 |
| Employment Status ($t$-1) 3 Migration | | | | | | |
| Not employed, long distance mover | | | -1.025* | 0.170 | -0.972* | 0.200 |
| Age group | | | | | | |
| 16–24 (base) | | | | | | |
| 25–34 | -0.263* | 0.083 | -0.182* | 0.066 | -0.157* | 0.077 |
| 35–44 | -0.314* | 0.094 | -0.221* | 0.071 | -0.183* | 0.085 |
| 45–54 | 0.042 | 0.103 | 0.063 | 0.076 | 0.133 | 0.092 |
| 55–64 | 1.256* | 0.111 | 0.806* | 0.081 | 1.059* | 0.098 |
| Qualifications | | | | | | |
| No qualification (base) | | | | | | |
| Intermediate qualification | -0.650* | 0.073 | -0.303* | 0.041 | -0.392* | 0.056 |
| Degree and higher | -1.016* | 0.110 | -0.453* | 0.063 | -0.689* | 0.087 |
| Marital status | | | | | | |
| Married (base) | | | | | | |
| Cohabiting | 0.027 | 0.068 | -0.008 | 0.047 | -0.034 | 0.058 |
| Car ownership | | | | | | |
| No car | | | | | | |
| 1 car | -0.491* | 0.078 | -0.291* | 0.057 | -0.356* | 0.069 |
| 21 cars | -0.740* | 0.084 | -0.453* | 0.060 | -0.525* | 0.073 |
| Youngest child | | | | | | |
| No children (base) | | | | | | |
| 0 – 4 years | 1.110* | 0.075 | 0.646* | 0.061 | 0.796* | 0.073 |
| 5 – 10 years | 0.249* | 0.080 | 0.127 | 0.065 | 0.163* | 0.078 |
| 11 – 15 years | -0.202* | 0.085 | -0.139 | 0.071 | -0.151 | 0.086 |
| Nondependent children | -0.322* | 0.085 | -0.134* | 0.068 | -0.177* | 0.082 |
| Household size | | | | | | |
| 2 or fewer people | | | | | | |
| 3 people | 0.401* | 0.076 | 0.207* | 0.060 | 0.246* | 0.073 |
| 4 or more people | 0.439* | 0.095 | 0.270* | 0.072 | 0.337* | 0.087 |
| Region | | | | | | |
| London (base) | | | | | | |
| Other metropolitan | 0.073 | 0.121 | 0.009 | 0.071 | 0.064 | 0.096 |
| Rest of South East | 0.157 | 0.104 | 0.065 | 0.062 | 0.125 | 0.085 |
| Rest of South and North | 0.077 | 0.109 | -0.002 | 0.065 | 0.066 | 0.088 |
| Employment rate | | | | | | |
| Women's rate , 75 percentile men's (base) | | | | | | |
| Women's rate 2: 75 percentile men's | -0.254* | 0.067 | -0.161* | 0.040 | -0.205* | 0.054 |
| Cambridge score ($t$-1) | | | | | | |
| Women's Cambridge score ::; men's (base) | | | | | | |

Table 4. *Continued*

| Categories | Model 1, Random-Effects Probit Parameters | Model 1 SE | Model 2, Random-Effects Probit Parameters | Model 2 SE | Model 3, Random-Effects Dynamic Probit[a] Parameters | Model 3 SE |
|---|---|---|---|---|---|---|
| Women's Cambridge score . men's | -0.905* | 0.036 | -0.302* | 0.033 | -0.247* | 0.038 |
| Constant | -0.021 | 0.156 | -0.822* | 0.107 | -0.797* | 0.135 |
| *k* | 0.681 | 0.011 | 0.249 | 0.018 | 0.429 | 0.023 |
| Log likelihood | -8,270.2 | | -7,542.4 | | -8,059.7 | |

*Note:* Number of person-waves ¼ 22,354, number of women ¼ 3,617.

[a]The random-effects and dynamic probit models involve different normalizations to the standard probit model (Stewart, 2005, 2006). For comparisons the former needs to be multiplied by an estimate of $r_u$=$r_v$ ¼ $\sqrt{1 - k}$.

*$p$ , .05.

The outcomes for women who were not employed at $t - 1$ displayed a different pattern. The probability of being out of work was similar for all groups not employed at time $t - 1$ except for those who moved for the woman's or both jobs (this group had a much lower probability of being out of work). Once again, these results were similar to those for the random-effects probit model (Table 5, Figure 3b), although the dynamic model results suggested a slightly reduced impact of state dependence. Comparing Panels b and c of Figure 3, we see that the broad patterns remain, although the probability of being out of work reduces slightly for women

FIGURE 2. A - C: PROBABILITY OF BEING NOT EMPLOYED (UNEMPLOYMENT OR ECONOMIC INACTIVITY; MODELS 1 – 3, TABLE 3).

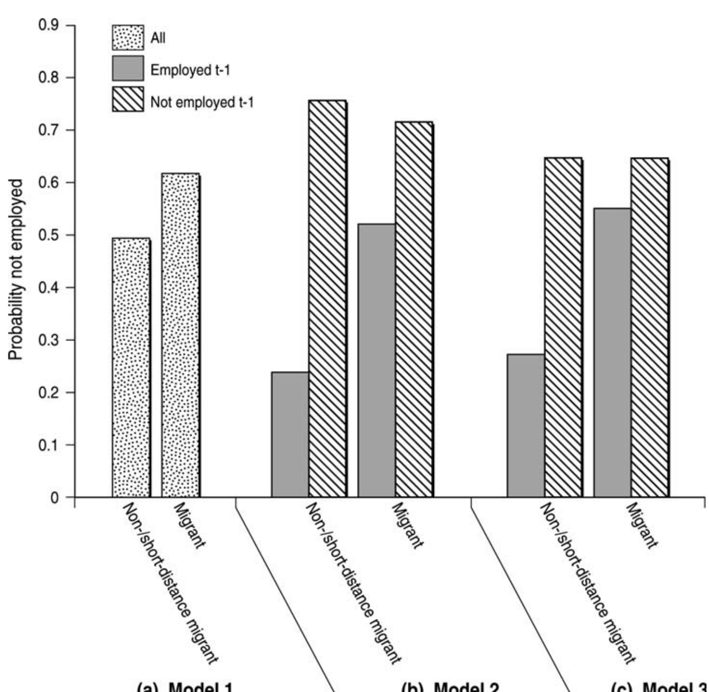


Table 5. *Model Parameters for Woman's Employment Status (y ¼ 1, Not Employed) Using Reason for the Move*

| Categories | Model 4, Random-Effects Probit | | Model 5, Random-Effects Probit | | Model 6, Random-Effects Dynamic Probit[a] | |
|---|---|---|---|---|---|---|
| | Parameters | *SE* | Parameters | *SE* | Parameters | *SE* |
| Employment status ($t$-1) | | | | | | |
| Employed (base) | | | | | | |
| Not employed | | | 1.674* | 0.038 | 1.341* | 0.044 |
| Reason for move | | | | | | |
| Nonmigrant (base) | | | | | | |
| Moved woman's or both jobs | -0.397* | 0.192 | 0.176 | 0.201 | 0.325 | 0.223 |
| Moved man's job | 0.665* | 0.146 | 1.030* | 0.150 | 1.088* | 0.163 |
| Moved other reasons | 0.212* | 0.053 | 0.389* | 0.058 | 0.380* | 0.067 |
| Employment status ($t$-1) 3 Reason for move | | | | | | |
| Not employed, moved woman's or both jobs | | | -1.517* | 0.329 | -1.435* | 0.361 |
| Not employed, moved man's job | | | -1.198* | 0.249 | -1.203* | 0.272 |
| Not employed, moved other reasons | | | -0.487* | 0.090 | -0.441* | 0.105 |
| Age group | | | | | | |
| 16 – 24 (base) | | | | | | |
| 25 – 34 | -0.251* | 0.083 | -0.179* | 0.066 | -0.151* | 0.077 |
| 35 – 44 | -0.299* | 0.094 | -0.212* | 0.071 | -0.171* | 0.085 |
| 45 – 54 | 0.059 | 0.103 | 0.073 | 0.076 | 0.148 | 0.092 |
| 55 – 64 | 1.273* | 0.111 | 0.811* | 0.081 | 1.067* | 0.098 |
| Qualifications | | | | | | |
| No qualification (base) | | | | | | |
| Intermediate qualification | -0.650* | 0.073 | -0.303* | 0.041 | -0.390* | 0.056 |
| Degree and higher | -1.010* | 0.110 | -0.440* | 0.063 | -0.670* | 0.086 |
| Marital status | | | | | | |
| Married (base) | | | | | | |
| Cohabiting | 0.009 | 0.068 | -0.021 | 0.047 | -0.043 | 0.058 |
| Car ownership | | | | | | |
| No car | | | | | | |
| 1 car | -0.490* | 0.078 | -0.292* | 0.056 | -0.354* | 0.068 |
| 21 cars | -0.738* | 0.084 | -0.457* | 0.059 | -0.525* | 0.073 |
| Youngest child | | | | | | |
| No children (base) | | | | | | |
| 0 – 4 years | 1.115* | 0.075 | 0.647* | 0.061 | 0.793* | 0.073 |
| 5 – 10 years | 0.257* | 0.080 | 0.133* | 0.065 | 0.165* | 0.078 |
| 11 – 15 years | -0.195* | 0.086 | -0.130 | 0.071 | -0.150 | 0.086 |
| Nondependent children | -0.314* | 0.085 | -0.125 | 0.068 | -0.170* | 0.082 |
| Household size | | | | | | |
| 2 or fewer people | | | | | | |
| 3 people | 0.387* | 0.076 | 0.199* | 0.060 | 0.242* | 0.072 |
| 4 or more people | 0.429* | 0.095 | 0.266* | 0.071 | 0.337* | 0.087 |
| Region | | | | | | |
| London (base) | | | | | | |
| Other metropolitan | 0.097 | 0.120 | 0.015 | 0.070 | 0.073 | 0.096 |
| Rest of South East | 0.181 | 0.104 | 0.073 | 0.062 | 0.139 | 0.084 |
| Rest of South and North | 0.104 | 0.109 | 0.004 | 0.064 | 0.077 | 0.087 |
| Employment rate | | | | | | |

Table 5. *Continued*

| Categories | Model 4, Random-Effects Probit | | Model 5, Random-Effects Probit | | Model 6, Random-Effects Dynamic Probit[a] | |
|---|---|---|---|---|---|---|
| | Parameters | *SE* | Parameters | *SE* | Parameters | *SE* |
| Women's rate ≤ 75 percentile men's (base) | | | | | | |
| Women's rate ≥ 75 percentile men's | -0.241* | 0.067 | -0.156* | 0.040 | -0.199* | 0.053 |
| Cambridge score ($t-1$) | | | | | | |
| Women's Cambridge score ≤ men's (base) | | | | | | |
| Women's Cambridge score > men's | -0.910* | 0.036 | -0.304* | 0.033 | -0.251* | 0.038 |
| Constant | -0.063 | 0.156 | -0.857* | 0.107 | -0.842* | 0.134 |
| $k$ | 0.681 | 0.011 | 0.243 | 0.018 | 0.423 | 0.023 |
| Log likelihood | -8,264.5 | | -7,516.7 | | -8,046.0 | |

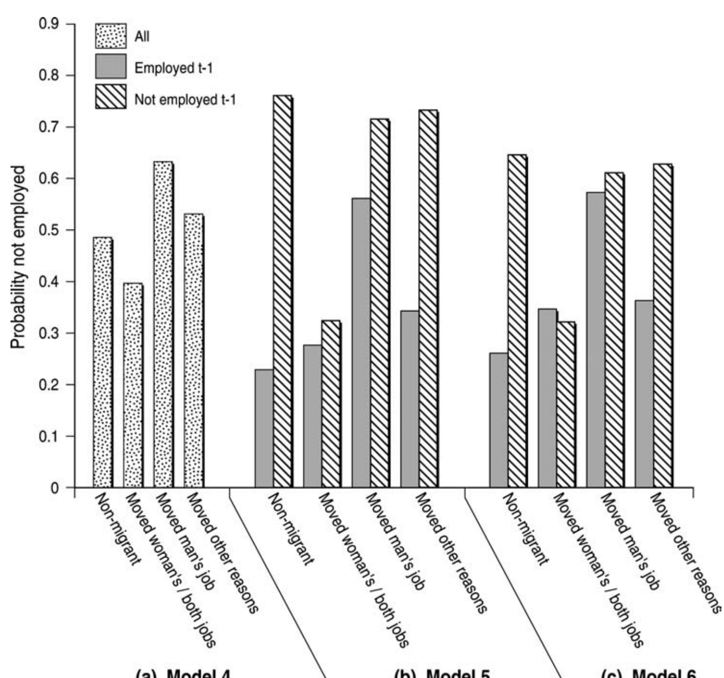
[a] The random-effects and dynamic probit models involve different normalizations to the standard probit model (Stewart, 2005, 2006). For comparisons the former needs to be multiplied by an estimate of $r_u/r_v = \sqrt{1 - k}$.

*$p < .05$.

who were not employed at $t - 1$ and reduces quite considerably for women who moved for their own or both jobs. For women who were employed at $t - 1$, however, moving for any reason increases the probability of being out of work, with this effect being most pronounced (and significant) for those who moved for the man's job.

The results for the other explanatory variables are consistent with those described for Model 3

FIGURE 3. A - C: PROBABILITY OF BEING NOT EMPLOYED (UNEMPLOYMENT OR ECONOMIC INACTIVITY; MODELS 4 - 6, TABLE 4).



(Table 4). The oldest group (ages 55 – 64) and those with young children were significantly more likely to be out of work, and those with higher qualifications, who owned one or more cars, had nondependent children, were living in areas where women's employment rates were high, and had a higher male Cambridge Scale score than their partner were less likely to be out of work.

To test whether attrition bias influenced our results we refitted our six models for the 1,043 women in our sample who appeared in all waves and answered all the relevant questions. Although some parameters changed slightly, these related mainly to the control variables in our models. The parameters for our key variables (migration and employment at $t - 1$) remained virtually identical in direction, size, and significance. Overall, the sensitivity analysis suggested that our modeling results based on the full data set were robust.

## DISCUSSION

The longitudinal analysis presented here provides some important new insights into the relationship between family migration and employment status. Unlike virtually all previous family migration studies, which focus on the distance moved or migration across some arbitrary administrative boundary, we have considered the reason for moving. This allows us to test the family migration hypothesis directly by assessing the impact on women of moving for the man's job, rather than using distance as a surrogate for employment-related moves, most of which are assumed to be on behalf of the man's job. We also compared the results with those based on a distance-based measure. In many cases, long-distance moves will be for employment reasons, but there are quite a number of such moves that occur over shorter distances. Similarly, a certain proportion of moves for nonemployment reasons, particularly associated with education for young adults and family-related factors for older people, will be over long distances. Second, we investigated the potential importance of state dependence by including a lag of the *y* variable (employment status) in the analysis. The importance of this issue is well recognized in the employment literature within economics, but it seems to have been ignored in the family migration literature, except in a recent study by Taylor (2007).

The descriptive results based on the reason for moving demonstrate some of the problems of using a distance moved variable. Many moves over 30 km are for nonemployment-related reasons, whereas a considerable number of short distance moves are for employment reasons, especially if the move is associated with the woman's job. We also found that there is significant state dependence, that is, the women's current employment status is infl ed by her employment status in the previous year. Ignoring the effects of previous employment could potentially result in biased estimates. Previously employed women who move a long distance are signifi y more likely to be out of work following the move, whereas moving results in a slight benefit for women who were previously out of work.

Focusing on the reason for the move in the modeling results, the negative effect of moving on previously employed women was caused mainly by moves associated with the man's job, which is consistent with the central hypothesis advanced within the family migration literature. The probability of being out of work following a move for the man's job was well over twice that of nonmigrants for previously employed women. We also found, however, that previously employed women who move for other reasons were also significantly more likely to be out of work than nonmigrants. This group will include movers for family, environmental, and housing reasons, and it suggests that the burden of migration may not only be associated purely with moves on behalf of the man's job. As we argued above, migration in general may have negative impacts on women in a number of ways, and this has also been shown to be the case in relation to migration-related union dissolution (see Boyle, Kulu, Cooke, Gayle, & Mulder, 2008). We also note that previously employed women who moved for their job or for both jobs were no more likely to be employed at time *t* than nonmigrant women. Although moving for this group may be beneficial for other reasons, it does not have a positive effect on their own employment status.

For the women who were not working at time $t - 1$, moving for the woman's or both jobs reduced the probability of being out of work to less than half that of nonmigrant women. Moving for the man's job or for other reasons lowered their chances of being out of work slightly, but this effect was not significant.

Overall, our results are important, as they demonstrate that previously employed women do suffer from family migration. Many previous studies have compared women's employment status after the move and have failed to explore whether women who moved were more likely to be out of work prior to the move. One possibility is that women appear more likely to be out of work following family migration, but in fact the effect is driven by the movement of women who were previously not employed. Our results suggest that it is not the case that women who were previously out of work explain the apparent association between family migration and women being out of work. Thus, our results provide convincing evidence of a family migration effect, but demonstrate that state dependence is a crucial factor. We also show that using a simple distance moved measure to distinguish probable employment-related moves is rather simplistic, as some long-distance moves have a much greater effect on women's employment status than others, depending on the reason for the move. In particular, moves for the man's job for women who were employed previously had a pronounced effect (note that 50% of those who moving for the man's job moved less than 30 km). These results provide an important, and previously underexplored, contribution to the family migration literature.

## REFERENCES


Arulampalam, W. (2001). Is unemployment really scarring? Effects of unemployment experiences on wages. *Economic Journal*, *111*, F585 – F606.

Arulampalam, W., Booth, A. L., & Taylor, M. P. (2000). Unemployment persistence. *Oxford Economic Papers*, *52*, 24 – 50.

Arulampalam, W., Gregg, P., & Gregory, M. (2001). Introduction: Unemployment scarring. *Economic Journal*, *111*, F577 – F584.

Bailey, A. J., & Cooke, T. J. (1998). Family migration and employment: The importance of migration history and gender. *International Regional Science Review*, *21*, 99 – 118.

Becker, G. S. (1974). A theory of social interactions. *Journal of Political Economy*, *82*, 1063 – 1093.

Bielby, W. T., & Bielby, D. D. (1992). I will follow him: Family ties, gender-role beliefs, and reluctance to relocate for a better job. *American Journal of Sociology*, *97*, 1241 – 1267.

Böheim, R., & Taylor, M. (2002). Tied down or room to move? Investigating the relationships between housing tenure, employment status and residential mobility in Britain. *Scottish Journal of Political Economy*, *49*, 369 – 392.

Boyle, P. J., Cooke, T. J. Halfacree, K. H., & Smith, D. (1999a). Gender inequality in employment status following family migration in GB and the US: The effect of relative occupational status. *International Journal of Sociology and Social Policy*, *19*, 115 – 150.

Boyle, P. J., Cooke, T. Halfacree, K. H., & Smith, D. (1999b). Integrating GB and US census microdata for studying the impact of family migration on partnered women's labour market status. *International Journal of Population Geography*, *5*, 157 – 178.

Boyle, P. J., Cooke, T. J., Halfacree, K. H., & Smith, D. (2001). A cross-national comparison of the impact of family migration on women's employment status. *Demography*, *38*, 201 – 213.

Boyle, P. J., Cooke, T. J., Halfacree, K. H., & Smith, D. (2002). A cross-national comparison of the effects of family migration on women's labour market status: Some diffi lties with integrating microdata from two censuses. *Journal of the Royal Statistical Society Series A (Statistics in Society)*, *165*, 465 – 480.

Boyle, P. J., Cooke, T. J., Halfacree, K. H., & Smith, D. (2003). The effect of long distance family migration and motherhood on partnered women's labour market activities in GB and the US. *Environment and Planning A*, *35*, 2097 – 2114.

Boyle, P. J., Halfacree, K. H., & Smith, D. (1999). Family migration and female participation in the labour market: Moving beyond individual-level analyses. In P. J. Boyle & K. H. Halfacree (Eds.), *Migration and gender in developed countries* (pp. 114 – 135). London: Routledge.

Boyle, P. J., Kulu, H., Cooke, T. J., Gayle, V., & Mulder, C. H. (2008). The effect of moving on union dissolution. *Demography*, *45*, 209 – 222.

Brett, J. M. (1980). The effect of job transfer on employees and their families. In C. L. Cooper & R. Paynes (Eds.), *Current concerns in occupational stress* (pp. 99 – 136). New York: Wiley.

Chay, K. Y., & Hyslop, D. R. (2000). *Identification and estimation of dynamic binary response panel data models*. Working Paper, University of California, Berkeley.

Clark, W. A. V., & Huang, Y. (2006). Balancing move and work: Women's labour market exits and entries after family migration. *Population, Space and Place*, *12*, 31 – 44.

Clark, W. A. V., & Withers, S. D. (2002). Disentangling the interaction of migration, mobility, and labor-force participation. *Environment and Planning A*, *34*, 923 – 945.

Clark, W. A. V., & Withers, S. D. (2007). Family migration and mobility sequences in the United States: Spatial mobility in the context of the life course. *Demographic Research, 17,* 591 – 622.

Cooke, T. J. (2001). ‘‘Trailing wife’’ or ‘‘trailing mother’’? The effects of parental status on the relationship between family migration and the labour-market participation of married women. *Environment and Planning A, 33*, 419 – 430.

Cooke, T. J., Boyle, P. J., Couch, K., & Feijten, P. (in press). A longitudinal analysis of family migration and the gender gap in earnings in the United States and Great Britain. *Demography*.

Davies, R. B., Elias, P., & Penn, R. (1992). The relationship between a husband’s unemployment and his wife’s participation in the labour force. *Oxford Bulletin of Economics and Statistics, 54,* 145 – 171.

Eliason, M., & Storrie, D. (2006). Lasting or latent scars? Swedish evidence on the long-term effects of job displacement. *Journal of Labour Economics, 24,* 831 – 856.

Goodman, J. S., & Blum, T. C. (1996). Assessing the non-random sampling effects of subject attribution in longitudinal research. *Journal of Management, 22,* 627 – 652.

Gregg, P. (2001). The impact of youth unemployment on adult unemployment in the NCDS. *Economic Journal*, *111*, F626 – F653.

Gregg, P., & Tominey, E. (2004). *The wage scar from youth unemployment*. CMPO Working Paper Series No. 04/097, University of Bristol.

Halfacree, K. H. (1995). Household migration and the structuration of patriarchy: Evidence from the USA. *Progress in Human Geography*, *19*, 159 – 182.

Heckman, J. J. (1980). Does unemployment cause future unemployment? Defi itions, questions and answers from a continuous time model of heterogeneity and state dependence. *Economica*, *47*, 247 – 283.

Heckman, J. J. (1981). The incidental parameters problem and the problem of initial conditions in estimating a discrete time-discrete data stochastic process. In C. F. Manski & D. McFadden (Eds.), *Structural analysis of discrete data with econometric applications*. Cambridge, MA: MIT Press.

Heckman, J. J. (1991). Identifying the hand of the past: Distinguishing state dependence from heterogeneity. *American Economic Review*, *81*, 75 – 79.

Heckman, J. J. (2001). Micro data, heterogeneity and the evaluation of public policy: Nobel lecture. *Journal of Political Economy, 109,* 673 – 748.

Hyslop, D. R. (1999). State dependence, serial correlation and heterogeneity in intertemporal labor force participation of married women. *Econometrica*, *67*, 1255 – 1294.

Jacobsen, J. P., & Levin, L. M. (1997). Marriage and migration: Comparing gains and losses from migration for couples and singles. *Social Science Quarterly*, *78*, 688 – 709.

Jacobsen, J. P., & Levin, L. M. (2000). The effects of internal migration on the relative economic status of women and men. *Journal of Socioeconomics*, *29,* 291 – 304.

Juerges, H. (1998). Vocationally motivated migration behavior in double-income households. An empirical analysis using GSOEP-data. *Zeitschrift Fur Soziologie*, *27,* 358 – 372.

Juerges, H. (2006). Gender ideology, division of housework, and the geographic mobility of families. *Review of Economics of the Household*, *4*, 299 – 323.

Lichter, D. T. (1980). Household migration and the labor market position of married women. *Social Science Research*, *9,* 83 – 97.

Lichter, D. T. (1982). The migration of dual-worker families: Does the wife’s job matter? *Social Science Quarterly*, *63*, 48 – 57.

Long, L. H. (1974). Women’s labor force participation and the residential mobility of families. *Social Forces*, *52*, 342 – 348.

Magdol, L. (2002). Is moving gendered? The effects of residential mobility on the psyhological well-being of men and women. *Sex Roles*, *47*, 553 – 560.

Makowsky, P. P., Cook, A. S., Berger, P. S., & Powell, J. (1988). Women’s perceived stress and well-being following voluntary and involuntary relocation. *Lifestyles*, *9*, 111 – 122.

Mincer, J. (1978). Family migration decisions. *Journal of Political Economy*, *86,* 749 – 773.

Morrison, D. R., & Lichter, D. T. (1988). Family migration and female employment: The problem of underemployment among migrant married women. *Journal of Marriage and the Family*, *50*, 161 – 172.

Muhleisen, M., & Zimmermann, K. F. (1994). A panel analysis of job changes and unemployment. *European Economic Review*, *38*, 793 – 801.

Narendranathan, W., & Elias, P. (1993). Influences of past history on the incidence of youth unemployment: Empirical findings for the UK. *Oxford Bulletin of Economics and Statistics*, *55*, 161 – 185.

Prandy, K. (2000). The social interaction approach to the measurement and analysis of social stratification. *International Journal of Sociology and Social Policy*, *19*, 215 – 249.

Shannon, C. E. (1948). A mathematical theory of communication. *Bell System Technical Journal*, *27*, 379 – 423.

Shauman, K. A., & Noonan, M. C. (2007). Family migration and labor force outcomes: Sex differences in occupational context. *Social Forces*, *85*, 1735 – 1764.

Shihadeh, E. S. (1991). The prevalence of husband-centered migration: Employment consequences for married mothers. *Journal of Marriage and the Family, 53*, 432 – 444.

Sjaastad, L. A. (1962). The costs and returns of human migration. *The Journal of Political Economy*, *70*, 80 – 93.

Smits, J. (1999). Family migration and the labour-force participation of married women in the Netherlands, 1977 – 1996. *International Journal of Population Geography, 5*, 133 – 150.

Spitze, G. (1984). The effects of family migration on wives' employment: How long does it last? *Social Science Quarterly, 65*, 21 – 36.

Stewart, M. B. (2005). *The inter-related dynamics of unemployment and low-wage employment*. Warwick Economic Research Papers, Number 741, Department of Economics, University of Warwick.

Stewart, M. B. (2006). *-redprob- A Stata program for the Heckman estimator of the random effects dynamic probit model*. Mimeo, University of Warwick.

Taylor, M. (2007). Tied migration and subsequent employment: Evidence from couples in Britain. *Oxford Bulletin of Economics and Statistics, 69*, 795 – 818.

van den Berg, G. J., & van Ours, J. (1996). Unemployment dynamics and duration dependence. *Journal of Labour Economics*, *14*, 100 – 125.

Weissman, M. M., & Paykel, E. S. (1972). Moving and depression in women. *Society, 9*, 24 – 28.